An Asset Theory of Social Policy Preferences

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Abstract

This paper presents a theory of social policy preferences that emphasizes the composition of people's skills. The key to our argument is that individuals who have made risky investments in skills will demand insurance against the possible future loss of income from those investments. Because the transferability of skills is inversely related to their specificity, workers with specific skills face a potentially long spell of unemployment, or a significant decline in income, in the event of job loss. Workers deriving most of their income from specific skills therefore have strong incentives to support social policies that protect them against such uncertainty. This is not the case for general skills workers for whom the costs of social protection will weigh more prominently. We test the theory on public opinion data for 11 advanced democracies, and also show that differences in educational systems explain a substantial portion of cross-national variance in the level of social protection.

Introduction

It is a well known fact that human capital rivals physical capital as a source of personal and national wealth. Indeed, it is the single most important determinant of personal income in advanced industrialized countries. Yet, whereas physical assets -- buildings, machinery, goods, and money -- have long been recognized as essential for understanding the political interests of their owners, surprisingly little is known about the role of human capital in explaining public policy preferences. With the exception of trade policy,[1] it is only the *cognitive* aspects of education that have received systematic attention in explaining political preferences (Klingemann 1979; Kitschelt 1991; Duch and Taylor 1993).

Following Becker (1964), we conceptualize human skills as an *investment*, and ask how the character of this investment affects workers' preferences for social protection. We approach this question in a fashion that is similar to the way transaction cost economics explains the use of non-market institutions to overcome market failures (Williamson 1985). In a political version of this logic, endogenous trade theory hypothesizes that investments in physical assets that are specific to a particular location or economic transaction lead firms to lobby the state for protection against uninsurable risks (see Alt et al. 1999).[2] Since pulling out assets in response to adverse market conditions is difficult, firms will want protection against the effects of such conditions. We start from the similar idea that investment in skills that are specific to a particular firm or industry exposes their owners to risks for which they will seek non-market protection. Skills that are portable, by contrast, do not require extensive non-market protection, just as the exchange of homogeneous goods does not require elaborate non-market governance structures.[3]

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[1] The trade policy argument treats skilled labor as a more or less abundant factor, where a relatively abundant factor is hypothesized to favor free trade. See Rogowski (1987), Frieden (1991), and especially Scheve and Slaughter (1999) for different versions of this argument.

[2] Alt et al. (1999) show empirical evidence that lobbying rises with the asset specificity of industries. See also Alt et al. (1996) for a more theoretical treatment of this and related arguments concerning the importance of asset specificity.

[3] There is an analogy here to the argument that capital mobility increases the incentives of capital owners to oppose restrictions on trade and financial mobility. See Frieden (1991), and

Our theory does not necessarily contradict a long tradition in the study of the welfare state that emphasizes redistribution as a key political motive behind the welfare state (e.g., Korpi 1983; Esping-Andersen 1990). Indeed, Meltzer and Richard's (1981) influential median voter result for government spending, which focuses on the redistributive aspect of social protection, emerges as a special case in our model. Given a particular composition of skills, workers with higher income are likely to demand less social protection than workers with low income. Our argument parts ways with the Meltzer-Richard model, however, because we explicitly recognize that social protection also has an insurance aspect (Sinn 1995, Moene and Wallerstein 1999), and that demand for insurance varies between workers according to their degree of exposure to labor market risks (Baldwin 1992). Critically, in our model risk exposure is inversely related to the portability of workers' skills.

We test our model on public opinion data for 11 OECD countries and show that differences in individual support for social protection are indeed closely linked to whether workers derive their income primarily from general as opposed to specific skills. We also show that countries where the educational system favors investment in specific skills tend to have significantly higher levels of social protection than countries where the educational system favors investment in general skills. Furthermore, because specific skills countries tend to have more egalitarian wage structures, and because there is a redistributive element to social protection, our model helps explain why relatively egalitarian countries also tend to exhibit more popular support for redistributive spending. This link between equality and support for redistribution is an old puzzle in political economy that follows directly from the Meltzer-Richard model.

The remainder of this paper is divided into three sections. In the first we present the model and its main empirical implications. In the second section we test these implications on public opinion data from 11 OECD countries. The third section discusses the broader implications of the model for explaining differences in social protection across countries, while the concluding section pinpoints some topics for future research.

Bates, Brock, and Tiefenthaler (1991).

The Model

Assumptions

Workers derive their income from skills that can be either general or specific. Specific skills are skills that are valuable only to a single firm or to a group of firms (whether an industry or a sector), whereas general skills are portable across all firms. We distinguish between three different employment situations, or states of the world, each associated with distinct levels of income. In *State I* a worker is employed in a firm that utilizes both his specific and general skills; in *State II* the worker is employed in a firm that only utilizes his general skills; and in *State III* the worker is unemployed (i.e., none of his skills are being utilized).

In *State I* the worker is paid the total value of his specific and general skills, or *sg*, and in *State II* the worker is paid for his general skills, *g*, only. If a worker has no specific skills, then $s=1$ and she is always employed at the market value of her general skills. The key assumption is that general skills are marketable in all sectors of the economy, whereas specific skills are only marketable in one sector (the size of which is defined by the specificity of skills).

In addition to market income, workers receive transfer income from the government. In the neo-corporatist literature such income is sometimes referred to as a "social wage" (Cameron 1984), and it includes unemployment benefits, health care benefits, pensions, and other forms of non-wage compensation. We assume that transfers come in the form of a flat-rate payment, *R*, which incorporates the idea in the Meltzer-Richard model that there is a redistributive aspect to social protection.[4] Following the terminology in Estevez-Abe et al. (1999), one can distinguish between transfers that go to support the income of employed workers, *wage protection*, and transfers that go to the unemployed, *unemployment protection*.[5] In the development of the model we will discuss what happens if *R* only goes to unemployment protection. But in the main model

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[4] This is also a realistic assumption with after-tax income distributed significantly more equally than pre-tax income (see Gottschalk and Smeeding 2000 and Huber and Stephens 2001).

[5] A third type of protection identified by Estevez-Abe et al. (1999) is called *employment protection* and refers to legal and other barriers to layoffs. This type of protection could be modeled as a probability of keeping a job where a worker's skills fully utilized.

we will assume that all workers receive the same flat-rate subsidy, which may simply be referred to as *income protection*.

Transfers are paid out of a flat-rate tax (t) on all wages. Total per capita receipts are T, and all receipts are spent on transfers (i.e., we assume balanced budgets). As in the Meltzer-Richard model, taxation is assumed to create work disincentives, captured here by the following simple labor supply function:

$$(1) \qquad l(t) = 1/(1+t),$$

where $l(t)$ is the number of hours worked or the intensity of effort (the particular form of this function is chosen for mathematical convenience). This implies that total tax income is

$$(2) \qquad T = \frac{t \cdot w}{1+t} = R,$$

where $w/(1+t)$ is average taxable income.

Figure 1 illustrates the three states of labor market, and shows the disposable (after tax) income associated with each state: $I: \overline{sg}$, $III: R$, *and* $II: \overline{g}$.

[Figure 1 about here]

For a given period of time, there is a probability, p, of losing one's job, and another probability, q, of re-employment. In equilibrium $p \cdot e = q \cdot g$, where e is the share of employed workers and g is the share of the workforce that is unemployed $(e = 1 - g)$. This implies that in equilibrium $e=q/(p+q)$. Furthermore, if r is the probability that an employed worker is in *State I* (i.e., is in a job where both general and specific skills are utilized), then the equilibrium share of the labor force employed in *State I* is

$$(3) \qquad a = r \cdot e = r \cdot q/(p+q).$$

Likewise, the share of the labor force employed in *State II* is

$$(4) \qquad b = (1-r) \cdot e = (1-r) \cdot q/(p+q),$$

while the share of the labor force in *State III* (unemployment) is

(5) $g = p/(p+q)$.

For any individual worker with both specific and general skills, the proportions a , b , and g can be interpreted as probabilities in a lottery with three possible outcomes. An employed *sg*-worker will therefore seek to maximize the expected utility of income across all three states. Ignoring the discounting of future income (which makes no substantive difference to our results), this is captured by the following utility function:

(6) $V = a \cdot u(\bar{s}\bar{g}) + b \cdot u(\bar{g}) + g \cdot u(R)$,

where $u(.)$ is the worker's utility from net income, which for simplicity we assume is spent on consumption c. Using standard assumptions, we impose the following constraints on u:

$$\begin{aligned} & u_c > 0, \\ & u_{cc} < 0, \text{ and} \\ & \lim_{c \to 0} u'(c) = \infty. \end{aligned} \qquad (7)$$

A number of the results below hold for this general form of utility function (notably the Meltzer-Richard results). However, since the insurance function of the social wage will play an important role and since we then need specific conditions on risk aversion, we use the standard assumption of a constant Arrow-Pratt relative risk aversion (*RRA*) utility function. Specifically,

$$\begin{aligned} u(c) &= \frac{c^{1-a}}{1-a} \quad \forall\, a > 0, \neq 1 \\ &= \log c \quad \textit{for} \quad a = 1 \end{aligned} \qquad (7a)$$

With these assumptions in mind, we can now find workers' utility-maximizing preferences for social protection.

Optimizing social preferences

The logic of the presentation in this section is as follows. We first consider a simple base-line model with no insurance effects, no tax disincentives, only general skills, and no unemployment (section i). We then introduce tax disincentives to get the Meltzer-Richard result (section ii), and

subsequently add insurance effects (and unemployment) to get the Moene-Wallerstein result (section iii). Finally, we show what happens to the demand for social protection when the composition of skills is allowed to vary (section iv). To keep the presentation simple, all proofs are put in an appendix.

(i) *No insurance effects, no disincentive effects: the t=1 model.* In solving workers' maximization problem we begin by assuming a labor force with only general skills (s=1), and no unemployment (e=1). The simplest case is where there are no tax disincentive effects on the number of hours supplied (so that $l(t)=1$ rather than $l(t)=1/(1+t)$ as we shall subsequently assume).

When s=1, e=1, and $l(t)$=1 equation (6) reduces to:

$$V = u\big((1-t)g + tw\big) = u\big(g(1-R/w) + R\big) = \frac{1}{1-a}\cdot\big(g(1-R/w) + R\big)^{1-a} \tag{8}$$

We want to choose R to maximize V, where R is bounded between R=0, corresponding to t=0, and R=w, corresponding to t=1. Since

$$V_R = (g(1-R/w) + R)^{-a}\cdot(1-g/w)$$

and, since $0 \le R \le w$, $g>w$ implies that V_R is uniformly negative for all values of R and hence t: maximization of V therefore requires $t = 0$. Thus voters with above average income will choose a zero tax rate. And analogous argument for $g < w$ implies that voters with below average income will choose the maximum tax rate of 100%. This is the standard result that in the absence of insurance functions and tax disincentives, voters will want the maximum R (i.e., t=1) if $g<w$ and a zero R (t=0) if $g>w$. Thus, if the median voter, M, has an income less than the average income of w, the median voter will always vote for a maximum tax rate. The result is illustrated in Figure 2, *panel a*.

[Figure 2 about here]

(ii) *Disincentive effects, no insurance effects: the Meltzer-Richard model*. If we now include the tax disincentive effect that $l(t)=1/(1+t)$, we have:

$$(9)\qquad \begin{aligned} V &= u\left(\frac{1-t}{1+t}g+\frac{tw}{1+t}\right) = u\big(g(1-2R/w)+R\big) \\ &= \frac{1}{1-a}\cdot\big(g(1-2R/w)+R\big)^{1-a} \end{aligned}$$

This implies

$$(10)\qquad V_R = (g(1-2R/w)+R)^{-a}\cdot(1-2g/w)$$

so that in the Meltzer-Richard model, only voters with a g-level below that of half the average wage ($g=w/2$) will vote for a maximum tax rate. As illustrated in Figure 2, *panel b*, if the median voter has a g-level above $w/2$ he will therefore not vote for the maximum tax rate. Because of the simplicity of our tax disincentive function, voters with g levels below $w/2$ will vote for $t=1$ and voters with g levels above $w/2$ will vote for $t=0$.[6] With the more complex tax disincentive function used by Meltzer-Richard, workers with income in the range $[w/2, w]$ will prefer taxation up to the point where the benefits to them from redistribution are exactly outweighed by the costs of tax disincentives. If the median voter is in this range, as the Meltzer-Richard model assumes, then she or he may vote for a positive tax rate less than 1 (as illustrated in Figure 2).

One of the implications of the Meltzer-Richard model is that voter turnout will be related to the level of government transfers because non-voting tends to be concentrated among low-income people (Lijphart 1997). There is some cross-national evidence for this proposition (see Franzese 1998). On the other hand, there is little empirical support for another key implication of the model, developed by Alesina and Rodrik (1994), namely that relatively inegalitarian societies will exhibit greater pressures for redistributive spending than relatively egalitarian ones (see Perotti 1996 for a review of the evidence). Among advanced countries the relationship is actually the reverse (Bénabou 1996).

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[6] Meltzer-Richard have a more general tax disincentive function than that used here. In consequence the tax rate which maximizes tax revenue can be less than one.

(iii) *Disincentive effects, insurance effects: the Moene-Wallerstein model.* The Moene-Wallerstein model offers one possible explanation, which results from the introduction of insurance effects. For insurance effects to matter, we need at least two states of the world. In a simple version of the Moene-Wallerstein model (Moene and Wallerstein 1999), it is assumed that workers can either be employed at a gross wage equal to their "tax-incentivised" skill level $g/(1+t)$ or be unemployed. There are no specific skills (s=1), so equation (6) becomes:

(11) $$V = b \cdot u(\bar{g}) + g \cdot u(R)$$

Moene-Wallerstein show that if relative risk aversion is greater than unity, workers will choose a higher tax rate as they become wealthier: in other words, their aversion to risk outweighs the increased cost to them of insurance as their income increases.

To get the Moene-Wallerstein results on risk aversion, we assume that $\bar{g} = g \cdot (1 - 2R / w)$ so that R is only paid to those who are unemployed. It can then be shown (as we do in Appendix A) that

(12) $dR/dg < 0$ *iff* $RRA < 1$, and

(13) $dR/dg > 0$ *iff* $RRA > 1$.

Put differently, when risk-aversion is high (RRA>1), and if all transfers go to the unemployed, the relationship between income (g) and the preferred level of social protection is positive (see Figure 2, *panel c*). A key implication of this result is that, contrary to the Meltzer-Richard model, a means-preserving increase in inequality will *reduce* the median voter's preferred level of social protection (provided that the income distribution is skewed to the right). The reason is that such a rise in inequality lowers the income of the median voter, and since the insurance motive dominates the redistribution motive (RRA>1), demand for social protection will decline. In the Meltzer-Richard model there is no insurance motive, so a fall in the income of the median voter always leads to a rise in the demand for social protection. Risk aversion thus poses one potential solution to the empirical puzzle of why income equality is linked to spending.

Yet, despite the neatness of this result, our econometric estimations rather clearly reject its

implication that people prefer less redistribution at lower levels of income. This leaves the negative relationship between redistribution and inequality as an important unsolved puzzle for comparative political economy. We discuss below how our distinction between specific and general skills permits an alternative and more plausible interpretation.

(iv) *Disincentive effects, insurance effects, specific and general skills: the asset model.* This is the most general model and requires us to consider all three states in Figure 1. We therefore return to the present value of utility given by equation (6):

(6) $V = a \cdot u(\bar{s}\bar{g}) + b \cdot u(\bar{g}) + g \cdot u(R)$

where

$$\bar{s}\bar{g} = s \cdot g \cdot (1 - 2R / w) + R$$
$$\bar{g} = g \cdot (1 - 2R / w) + R$$

In addition, we now introduce an important variable, expected income before taxes and transfers, *y*. This is simply defined as:

$$y \equiv a \cdot sg + b \cdot g$$

We then proceed to ask, first, up to what value of *y* is the chosen *R* maximal, i.e. $t = 1$; second, under what *RRA* conditions does *R* fall or rise as *y* rises above this value; and, third, what happens to the choice of *R* as the balance of general and specific skills changes, holding *y* constant?

In the first result we show that a worker will only choose the maximum tax rate if $y \leq w/2$. Stated formally (the proof is in Appendix A):

Result I: *Given the assumptions of Model (iv),* $t = 1$ *iff* $y \leq \frac{w}{2}$

With this property established, we now consider what happens to *R* when *y* increases. As we show in Appendix A, this yields the following result:

Result II: *Given the assumptions of Model (iv), holding s constant and with y > w/2,*

$$\operatorname{sgn}\frac{\partial R}{\partial y} = \operatorname{sgn}\left[RRA(\bar{s}\bar{g}) - \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2} \right] .$$

What this equality says is that the direction of the relationship between *R* and income (the sign, sgn, of $\frac{\partial R}{\partial y}$) depends on the level of risk-aversion, just as in the simple Moene-Wallerstein model. However, for income to be positively related to support for spending the *RRA* requirement is more stringent ($RRA > \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2}$) than before (*RRA*>1). The reason is that *R* now goes to the employed as well as to the unemployed, and since employed workers in the Moene-Wallerstein model only have an insurance incentive in relationship to unemployment, *RRA* must be higher for the insurance motive to dominate the redistribution motive. This implication is also demonstrated by Moene and Wallerstein.

Now we come to the critical result which differentiates our approach from previous ones. Central to the argument of the paper is the proposition that an increase in specific skills relative to general skills, holding constant the level of expected income, implies an increase in preferred *R*: put broadly, workers with specific skills will prefer higher taxes and social protection than workers with general skills. The following result is also proved in Appendix A:

Result III: *Assuming a constant relative risk aversion utility function and RRA>0,* $\frac{\partial R}{\partial s} > 0$ *holding y constant.*

In other words, as *s* rises, the preferred level of *R* also rises. The intuition behind this key result is that workers with specific skills have more to fear if they lose their job than workers with general skills. This is because specific skill workers who are laid off face the risk of being re-employed in a sector where their skills are not needed. If this happens they will lose some of their previous income. General skill workers do not face this problem because they are always paid the value of their general skills. Hence, the more income derived from specific as opposed to general skills – that is, the higher the ratio *s*/*g* – the greater the demand for income protection (*R*). The logic is illustrated in Figure 2, *panel d*, and implies that the median voter support for social protection depends on the composition of his or her skills.

Summarizing the results in this section: with the simplest set of assumptions – only one state of the world (employment), only general skills, and no tax disincentives – the politics of social spending is all about redistribution (class politics if you will): those with a wage below the mean will want a maximum rate of taxation (t=1) whereas those above the mean will want zero taxation. If we add tax disincentives, however, the cost of redistribution may deter those low-income workers closest to the mean from demanding confiscatory taxation, and the median voter is likely to be among those workers. This is the Meltzer-Richard model.

When an unemployment state is added to the model, an entirely new motive enters into workers' calculations of their interests: insurance against loss of income. If workers are sufficiently risk-averse, and if all transfers go to the unemployed, rising income may in fact be associated with *higher* demand for protection since high-income workers have more to lose than low-income workers. This is the Moene-Wallerstein model. If some transfers go to the employed, however, the threshold of risk-aversion for which this relationship holds goes up since transfers to the employed only serve redistributive purposes.

Finally, when differences in the specificity of skills are introduced, which require at least two employment states (Sector I and II in our model), the insurance motive plays a crucial role *even* when workers are only moderately risk-averse ($0<RRA<1$) and *even* when transfers are distributed to both employed and unemployed workers. The reason is that employed workers risk losing the income from their specific skills, regardless of their exposure to unemployment. This coupling between skills and demand for insurance thus transforms the relationship between income and social policy preferences. The next section explores whether this proposition is supported at the *individual level* using empirical evidence from public opinion surveys in 11 advanced democracies. We subsequently test the implication at the *macro level* using a larger sample of 19 democracies.

Testing the Model

Statistical Model

In the last section we showed in our Model IV that the relationship between the "preferred" level

of *R* and the two exogenous variables *y* (expected income) and *s* (skill specificity) is given by the implicit equations:

$$(14) \qquad \begin{aligned} V_R(R, s, g) &= 0 \\ y &= a \cdot s \cdot g + b \cdot g \end{aligned}$$

And from (14) we derived Result II that

$$\frac{\partial R}{\partial y} < 0 \quad if \quad 0 < RRA < \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2}$$

and Result III

$$\frac{\partial R}{\partial s} > 0 \quad if \quad 0 < RRA.$$

We show in Appendix B that

$$R = K + \frac{\partial R}{\partial y} \cdot y + \frac{\partial R}{\partial s} \cdot s$$

is the first-order Taylor expansion of (14). Thus our regressions take the form

$$(15) \qquad R = k + a \cdot y + b \cdot s.$$

By implication if our estimate of *a* is significantly different from zero and negative, we can infer that $0 < RRA < \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2}$. If *b* is significantly different from zero and positive, $0 < RRA$, so that skill specificity increases the demand for social protection. This is our main argument and hypothesis.

More generally, Model IV encompasses Models I, II and III. Hence, we can test for these models as well. Model I (Meltzer-Richard without tax disincentives) implies that $a=b=0$. Model II (Meltzer-Richard with tax disincentives) implies that $a<0$ and $b=0$. And Model III (Moene-Wallerstein) implies that $a>0$ and $b=0$.

The data

We use individual-level data from 11 advanced democracies obtained from two sets of national mass surveys conducted under the auspices of the International Social Survey Program (ISSP), one in 1996 the other in 1997 (ISSP 1999; 2000).[7] These surveys offer by far the best individual-level data on skills and preferences for social protection. In the following two sections we describe the operationalization of the dependent and independent variables.

Dependent variables

The 1996 survey contains four spending questions that closely match our theoretical emphasis on income protection (*R*). Three of the four are used in a cluster of questions that asks whether the respondent would like to see more or less government spending on a) unemployment benefits, b) health care, and c) pensions.[8] Reflecting an assumption in the model, respondents were warned that more spending may require higher taxes. The fourth variable is based on a question that asks whether the respondent favors government spending on declining industries for the purpose of protecting jobs.[9] This question is as much about job security as it is about income security, but the two are obviously closely related, and we expect specific skills workers to be more concerned than general skills workers with keeping their present job and income. Moreover, although the respondent was not explicitly told about the potential costs of government subsidies, such

[7] The 11 countries are: Australia, Britain, Canada, France, Germany, Ireland, Italy, Japan, Netherlands, Norway, Sweden, and the United States. Japan, using in both ISSP surveys, could not be included because of missing data on a key occupational variable (explained below).

[8] The exact question wording is: “Listed below are various areas of government spending. Please show whether you would like to see more or less government spending in each area. Remember that if you say "much more", it might require a tax increase to pay for it.” The respondent is then presented with different spending areas and the following range of possible responses: “1. Spend much more; 2. Spend more; 3. Spend the same as now; 4. Spend less; 5. Spend much less; 8. Can’t choose, don’t know.”

[9] The exact question wording is: “Here are some things the government might do for the economy. Please show which actions you are in favor of and which you are against. Please tick one box in each line.” One of the actions is: “Support for declining industries to protect jobs: 1. Strongly in favor of; 2. In favor of; 3. Neither in favor of nor against; 4. Against; 5. Strongly against; 8. Can’t choose, don’t know; 9. NA, refused.”

subsidies are widely acknowledged to be problematic for economic efficiency.

The survey also asked people whether they favored more or less spending on "culture and the arts" and "the environment." These policy areas are clearly unrelated to social protection, but they are nevertheless relevant to our argument because general education is often argued to *increase* support for spending on "post-materialist" activities, whereas our theory says that it reduces support for spending in the social policy area (Kitschelt 1991; Duch and Taylor 1993). Since one might object that our findings for skills reflect general ideological opposition to government spending among those with long formal educations, it is useful to be able to show that the relationship between skills and support for spending varies by policy area.

To check that our six variables are indeed proxies for the two spending dimensions, we conducted a factor analysis (see Table 1). As expected, the analysis yields two, and only two, dimensions, with the social spending variables having high factor loadings on the first dimension, and the postmaterialist variables having high factor loadings on the second dimension. This pattern is repeated in every individual country, with little variation in the factor loadings. Consequently we created two simple additive indexes, one for social spending, the other for postmaterialist spending (both are standardized to vary between 1 and 5).[10]

[Table 1 about here]

Independent variables.

We use two different approaches to the measurement of skill specificity, reflecting different aspects of the theoretical model. The first is to classify workers' skills, or the skills required to perform certain jobs, according to their degree of specialization or specificity. This is an attempt to gauge *s* directly. The second starts from the model assumption that the difficulty of finding a job where one's skills are needed is proportional to their specificity. This is an attempt to gauge s indirectly through *rq* (the probability of re-employment into *state I*).

The first approach is based on the ILO's detailed classification of people's occupations:

[10] Alternatively we could have used the factor loadings directly to construct the two indexes. However, it makes little difference to the results.

the International Standard Classification of Occupations (ISCO-88). ISCO-88 classifies workers in "occupations" based on two criteria: the *level* of skills required for an occupation, and the *degree of specialization* of those skills. ISCO-88 distinguishes between four broad skill levels, which are a function of "the range and complexity of the tasks involved" and explicitly dependent on informal as well as formal training (ILO 1999, 6). Skill level thus corresponds to ($s+g$) in our model. All other distinctions between occupations are based on the specialization of skills required to carry out particular jobs, reflecting "the type of knowledge applied, tools and equipment used, materials worked on, or with, and the nature of the goods and services produced" (ILO 1999, 6). Guided by this logic, the subdivision of skills proceeds through four levels of aggregation until a high degree of skill homogeneity is reached within each group.[11] At the most disaggregated level, called the unit level, there are 390 occupational categories with highly specific job descriptions.[12] The occupation of every single respondent in the ISSP surveys was classified according to ISCO at either the most detailed or second most detailed level.[13]

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[11] We are not claiming that homogeneity is equivalent in every unit group. Yet, skills that are clearly distinct from one another are unlikely to be in the same group at the most disaggregated level, and major groups with a highly diverse skill structure therefore will tend to have more minor and unit groups.

[12] Unit group 3144, for example, represents "air traffic controllers", which is a member of the minor group "ship and aircraft controllers and technicians" -- itself one of five categories in the major group called "technicians and associate professionals".

[13] In some countries individuals were classified using an earlier version of ISCO (ISCO-68), but it is possible to translate the earlier version into the later with considerable consistency using a coding scheme developed by Harry Ganzeboom at Utrecht University (see Ganzeboom and Treiman 1996 and http://www.fss.uu.nl/soc/hg/ismf for details). The Swedish occupational classification is based on an amended version of an older edition of ISCO. *Statistiska Centralbyrån* (Statistics Sweden) provided us with a conversion table to translate these codes into ISCO-88 in reasonably consistent manner. Britain uses its own national classification system, but it is closely related to ISCO-88 and likewise uses skills as the basis for the classification. We received the British translation codes from *UK National Statistics*. The only problematic case is Italy where the few broad categories used in the 1996 ISSP survey are completely unrelated to the ISCO-88 categories. Instead we went back to an earlier 1990 ISSP study (ISSP 1993), which contains a somewhat more detailed occupational variable for Italy. Using this variable in conjunction with information on educational levels enabled us to map the Italian codes to the 1-digit ISCO-88 level in a fairly consistent manner. Yet, because of the lack of direct

We can exploit the skill-based hierarchical structure of ISCO-88 to capture skill specialization by comparing the share of unit groups in any higher-level class to the share of the workforce in that class. The logic is that the number of unit groups in any higher-level class will be a function of the size of the labor market segment captured by that class, and of the degree of skill specialization of occupations found in that particular labor market segment. For example, 10 percent of the workforce across our countries are classified as "plant and machine operators and assemblers" (major group 8), whereas this group accounts for 70 out of the 390 unit groups, or 17 percent of all unit groups. If occupations at the unit group level are, on average, equally homogeneous in terms of skills, the disproportionate share of unit groups in major group 8 will reflect a greater degree of specialization of skills found within that major group. By dividing the share of unit groups (.17) by the share of the labor force (.8), we can therefore generate a measure of the average skill specialization within that particular major group (2.1). This calculation can also be carried out at the lower sub-major level, and we have used the mean of these calculations to get another proxy for *s*.[14] The resulting variable has 27 values ranging from 0.4-4.7.

Because the theoretical concept of skill specificity is a *relative* variable, the final step is to divide the absolute skill specialization measure, *s*, by the ISCO measure of the *level* of skills.[15] This gives us a proxy for $s/(s+g)$ that we will refer to as *s*_1. Alternatively we can divide *s* by a proxy for peoples' general skills, *g*, which gives us a measure for s/g. We call this alternative measure *s*_2. The proxy for *g* that we use is the respondent's highest academic degree as recorded

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correspondence, the results for Italy must be viewed with caution.

[14] The sensitivity of *s* to small differences in the number of unit groups assigned to each higher-level group is greater at lower levels of aggregation, and these differences may not accurately reflect differences in skill specificity. This source of error is minimized at the highest level of aggregation. However, the greater variance of the measure at lower levels of aggregation helps reduce the standard error on the estimated parameter for the skill variable.

[15] Using an absolute measure of s generates results that are downward biased. At the limit, if the (unknown) correlation between *s* and *g* is 1, *s* will have no effect on preferences. It is therefore important to develop relative measures.

by the respondent. The variable has five levels: 1. completed primary degree or lower[16]; 2. incomplete secondary; 3. completed secondary; 4. incomplete and completed semi-higher degree, or incomplete university degree; and 5. completed university degree.[17]

The second approach to measuring skill specificity is based on the observation that the probability of moving from any particular job into one that makes use of a worker's skills *(state I*) is *rq* for specific skills workers and *q* for general skills workers, where $r<1$. If we conceive of *rq* as an element in a continuum [0, *q*], *r* would then be a measure of the asset-specificity of a worker's skills. At the heart of the concept of job specificity is the idea that outside options are more limited for workers with specific skills than for workers with general skills.

The 1997 ISSP survey contains a question is that precisely taps workers' assessment of their outside options. The question reads as follows:

> If you were looking actively, how easy or difficult do you think it would be for you to find an acceptable job?

The respondent could answer "very easy", "fairly easy", "neither easy nor difficult", "fairly difficult", and "very difficult." The key here is that the difficulty of finding an acceptable job is likely to be related to how portable a person's skills are. High skill specificity means that there are fewer jobs where these skills are used, and the number of job openings is also likely to be smaller because asset-specific investments by employers and employees tend to lengthen tenure and limit turnover. In addition, the probability of finding an appropriate job close to a person's current residence, which is also a likely component of what an individual considers "acceptable," falls with the number of job openings in a given geographical area.[18] Asking people about the

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[16] In some countries a distinction is made between incomplete and complete primary education.

[17] Alternatively we could use years of formal schooling as a measure of *g*, but the results are very similar.

[18] Scheve and Slaughter (2001) argue, and show empirically, that home ownership can be treated as a relatively immovable asset that affects people's preferences for trade protection. It

probability of finding an acceptable job is therefore likely to generate answers that are systematically related to a person's skills. In the absence of extensive information about individual work histories, and employment conditions in particular labor market niches, the question is therefore about as good a measure of *rq* as one could hope for. We refer to it as *s*_3.

There is however an ambiguity in the relationship of *s*_3 to the theoretical concept of *s*. The reason is that we cannot know for sure if peoples' responses reflect their *absolute* level of specific skills or the *relative* share of their skills that is specific. To make sure that the skill measure is a relative measure we can divide *s*_3 by *g*. We call this alternative measure *s*_4. If *s*_3 is already a relative measure, we simply get another relative measure that should also be positively related to preferences for social spending.

The different skill measures are summarized in Table 2. Not surprisingly, the correlations are higher between measures using either the survey question *or* the ISCO classification. The lowest inter-correlations are between *s*_3 and *s*_1 or *s*_2. To some extent this may reflect that *s*_3 is an absolute rather than a relative measure, but the main reason is simply that there are more values, and therefore variance, on *s*_3 than on either *s*_1 or *s*_2. If we therefore use the average of *s*_3 for each category of *s*_1 and *s*_2, the inter-correlations jump to between .38 and .53. A similar effect occurs for *s*_4. To facilitate comparison of the effects of the different proxies in the subsequent regression analysis, all proxies for *s* have been divided by their standard deviation.

[Table 2 about here]

One final methodological issue needs to be addressed. Because the question used as the basis for *s*_3 and *s*_4 was asked only in the 1997 survey, whereas all the questions about spending were asked only in the 1996 survey, it was necessary to "translate" the 1997 information on *s*_3 so it could be used in the 1996 survey (*s*_4 can be always be calculated from *s*_3). For this purpose we calculated averages for *s*_3 at the 3-digit ISCO-88 level in the 1997, and then assigned these

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would be interesting to interact home ownership with the question about the difficulty of finding an acceptable job, but residential status is unfortunately not recorded by ISSP.

values to individuals in the 1996 survey based on their 3-digit ISCO classification in that survey.[19] Since the classification of occupations is motivated by the skills required in these occupations, it is reasonable to expect the original information about *s* is preserved to considerable extent in this translation. Moreover, because the 1996 and 1997 samples are drawn from the same populations,[20] we show in Appendix C that *s*_3, averaged by ISCO level-3 groups, is an unbiased estimator for the original variable.

In addition to the skill variables (*s*_1-4) we used self-reported pre-tax and transfer income as a proxy for *y*, and the following set of controls:

Age. Older workers are likely to be more concerned with job security and income than younger workers since their time to retirement is shorter and since their ability to find new employment is likely to be more limited.

Gender. As argued by Orloff (1993) and Estevez-Abe (1999), women may demand more protection than men in comparable jobs because they need to be able to leave, *and return*, to the labor market for the purpose of child rearing.

Union membership. Since one of the main functions of unions is to insure their members against labor market risks, it is reasonable to expect that union members are particularly concerned with social protection (see, for example, Korpi 1989).

Part-time employment. Part-time employees are often in vulnerable labor market positions, and this may cause particular concern for job security and income protection. On the other hand, part-time employees depend more on flexible labor markets to generate non-standard jobs, which suggests a countervailing effect.

Non-employed. Esping-Andersen (1999) has argued that some outsider groups may share an interest in social and economic policies that maximize their ability to enter employment. But this is an extremely heterogeneous group that may not have common policy preferences. We need to include the variable to control for the possibility that the non-employed have very different attitudes than the employed.

Unemployed. The expectation is obviously that the unemployed, relying as they do on transfers,

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[19] There are 116 unique groups at the 3-digit level. The more fine-grained 4-digit level is not available for some countries, and contains a large number of empty categories where it is.

[20] With the minor qualification that those who turned 18 between the 1996 and 1997 surveys were not part of the population in the former survey.

will support high levels of income protection.

Self-employment. The self employed are expected to favor free markets and low levels of social protection because the self-employed depend on flexible labor markets and often on relatively low-paid workers.

Information. It is conceivable that better information about the economy yields particular views on the desirability of social spending. There was an intense public debate about the proper role of the state in the 1990s, and it could be argued that better informed people would reflect the predominant view in this debate, which tended to see cut-backs as necessary on efficiency grounds (corresponding to a higher cost of distortionary taxation in our model).[21] Information is measured by a question that asked people to declare their degree of agreement with the following statement: "I feel that I have a pretty good understanding of the important political issues facing our country."[22]

Left-right position. Attitudes to social protection may in part be a reflection of people's ideological predispositions, or perhaps the socializing effects of political parties.[23] We control for this possibility by including positions on a left-right scale based on the respondent's support of parties that are ranked from far left to far right.

Findings

For each of the 11 countries, and for each of the three dependent variables, we estimated the regression model in equation (15), including a full set of controls. To cope with problems of missing observations we used a multiple imputation technique developed by Honaker et al. (1999). This strategy is superior to the traditional approach of "listwise deletion", which is both inefficient and potentially biased (King et al 2000).[24] The following presentation is divided into a section with the key results, and a section where we test the robustness of these results and

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[21] This idea was suggested by an anonymous reviewer.

[22] Respondents could answer: 1. Strongly agree; 2. Agree; 3. Neither agree nor disagree; 4. Disagree; 5. Strongly disagree. The variable was reversed so that higher values measure more information.

[23] This was suggested by an anonymous reviewer. We note that party support may in part be endogenous to skills. If so, the effect of skills will be underestimated by the parameter for *s*.

[24] In practice, however, our results are very similar to those obtained by using listwise deletion. The effects of our theoretical variables tend to be slightly stronger when we use listwise deletion, but the standard errors are also larger.

discuss potential objections to the way we interpret them.

The basic results

To give a sense of the central tendency of the estimates, Table 3 shows the results from a pooled analysis, including a full set of country dummies. Since the Italian survey was conducted in 1990 and lacks information on several of the control variables, it was not included in the calculation of these pooled results.[25] In the next section we show that the results for Italy are consistent with those presented in Table 3.

[Table 3 about here]

The model in column (1) uses the average of the four measures of skills, called *s*_composite, as a summary variable for skill composition. The next four columns show the results for each of the component measures (*s*_1-4). The parameter estimates for the controls are not shown in these columns because they are almost identical to those in (1). Model (6) is identical to (1) except that the regression now includes union membership as an independent variable. Since union membership was not recorded in Australia, the estimation of model (6) excludes this country.

In interpreting the results, first note that the parameters for income, *y*, and the four measures of skill, *s*_1-4, are in the predicted direction, and highly statistically significant. The negative effect of income implies that people's risk aversion is not sufficiently high to make their demand for transfers rise with income. Technically speaking, $RRA < \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2}$, which means that the Meltzer-Richard redistribution logic dominates the Moene-Wallerstein insurance logic.

Yet, for our purposes the key finding is the positive effect of specific skills on preferences for spending (which implies that $RRA>0$). Each of the four (standardized) skill variables is associated with significantly higher support for spending, and three of the four measures exhibit similar magnitudes of effects. The parameter for *s*_3 is lower than for the other measures, but this

[25] This is a purely presentational issue. Regressing spending preferences against the smaller set of independent variables that allows us to include Italy has negligible effects on the results for the income and skill variables. But it would require another column of data.

is not entirely unexpected given that this variable may capture absolute rather than relative endowments of specific skills (or a combination of absolute and relative endowments). For all correlations between *s* and *g* that are greater than -1, absolute measures of *s* will yield lower parameter estimates than relative measures.[26]

Considering the very different approaches to measuring skills, it is reassuring that the results are consistent across definitions. Yet, statistically significant effects do not necessarily imply large substantive effects. In Table 4 we have therefore estimated the portion of the explained variance accounted for by each of the independent variables, as well as the impact on preferences of a one standard deviation change in each of the independent variables.[27] The estimates are based on the results of model (6) in Table 3, which includes all the relevant variables.

[Table 4 about here]

Although we cannot precisely attribute the proportion of explained variance to each of the independent variables, we can calculate likely ranges. The *upper* bounds of these ranges are found by recording the increase in explained variance (measured as a percentage of the total explained variance) when a variable is included as the *first* predictor (apart from the country dummies). This number encompasses every direct, indirect, and spurious effect of the variable. The *lower* bounds are calculated as the increase in explained variance (as a percent of the total explained variance) when a variable is entered as the *last* predictor. This procedure eliminates all spurious effects of the variable, but also discounts all possible indirect effects. The true explanatory power of any variable is likely to be somewhere in between these bounds.

Using this method, Table 4 shows that income and skills are unambiguously the most important variables in explaining social policy preferences. Thus, income accounts for between 16

[26] In fact, the correlation between *s*_3 and a measure of *g* based on general education is close to 0 in our data, which implies an estimated effect of *s*_3 is half the "true" effect of skills.

[27] The fully specified model explains about 21 percent of the total variation in individual preferences, which is at the high end of the range for most analyses of public opinion data.

and 53 percent of the total explained variance, whereas skills account for between 19 and 37 percent. Jointly, income and skills capture between 48 and 73 percent of the explained variance, with the remainder accounted for by the controls.

The key role of income and skills in explaining social policy preferences is confirmed when we consider the impact of a one standard deviation change in these variables (column 3). A standard deviation change in either variable is associated with about 20 percent of a standard deviation change in preferences. Together, the impact of income and skills is as great as the joint effect of a standard deviation change in all controls simultaneously. Note also that the effects of both variables are estimated very precisely, varying in a narrow range between 0.12 and 0.14 (95-percent confidence interval).

The results for the controls also generally confirm our expectations. Individuals who are particularly exposed labor market risks – the unemployed, women, and older workers – are more favorably disposed to increasing social spending than others. The same is the case for union members, whereas the self-employed are more likely to oppose social spending. Those who consider themselves well-informed about politics are also more likely to oppose spending, perhaps reflecting a political reality at the time that was hostile to the welfare state. Supporters of right parties, not surprisingly, also express less support for social spending than supporters of left parties. Finally, we note that the attitudes of part-time employees and those outside the labor market are indistinct from others' attitudes. These groups are evidently too heterogenerous to share any common interest in social policies.

Gender stands out among the control variables, accounting for between 6 and 17 percent of the total explained variance, and having the greatest impact among the controls. As argued by Estevez-Abe (1999), women require more protection than men in comparable jobs because they need to be able to leave, *and return*, to the labor market for the purpose of child rearing. Yet, almost half of this effect disappears if the skill variable is removed from the equation. The reason, we believe, is closely related to our theoretical argument. Since women know that they are likely to leave their jobs before they can reap the full returns on specific skill investments, they are dissuaded from making such investments (Estevez-Abe et al. 1999). This shows up in our data as a negative effect of (female) gender on *s*. In other words, *if* women invest in specific skills they

are more prone than men to support high levels of social protection, but they are somewhat less likely to invest in these skills in the first place.

Robustness tests

In this section we test the robustness of the results, and address some potential objections to our interpretation of the results. We first note that the findings for *y* and *s* stand up to any combination of the controls included above, and they are robust to the inclusion of any other variable used in the survey, hereunder region, public sector employment, urbanization, and supervisory position -- in any combination.[28] But while income and skills are powerful explanatory variables in the pooled analysis, pooling can disguise considerable cross-national variation in the strength of the results, and sometimes estimated parameters can even reverse in particular cases. In addition, pooling usually yields exaggerated t-scores compared to those found for individual countries.[29] We therefore ran our regressions on each of the 11 countries individually. The results for the theoretical variables are shown in Table 5.

[Table 5 about here]

Note that every regression yields results that are consistent with the pooled analysis, with each of the 60 parameters recording the correct sign, and most being significant at the .01 level or better. The composite skill variable (bolded) is always significant at a .01 level or better, and for 9 of the 11 countries the parameter estimates for *s* vary in a fairly narrow range between 0.10-0.17 (the parameter in the pooled analysis is .14). Only Ireland and Italy fall slightly out of the pattern with parameters around .07. Yet, the effects for these countries are still statistically highly significant,

[28] None of these variables were used in every survey, so instead of cluttering the presentation with several additional columns, we left these variables out of the main analysis.

[29] The reason is that the standard error has the form (s.e. of equation error)/(s.e. of variable). Since the denominator is the square root of the sum of squares of the explanatory variable divided by *N*, this normally increases with *N* since a squared term is added on the top and 1 is added to the bottom (though it does not have to be so).

and it should be noted that *s*_comp in both cases are based on only two proxies for *s*. In the case of Italy, these proxies also use an occupational variable that maps rather poorly onto ISCO-88, potentially diluting the skill distinctions between categories.

As in the case of the pooled analysis, we also note that the results for *s*_3 are somewhat weaker across all cases than for the other skill measures, but only in one instance (the US) do we get a statistically insignificant result. Given the variety of countries and the differences in measurements, the combination of results is quite remarkable support for our theory.

An objection that can be raised is that we may in part be capturing an ideological aversion to government spending among those with higher education. Two of the measures of *s* have formal education in the denominator, and the other two implicitly assume that such skills are part of the denominator. In quantitative terms, general education accounts for roughly one third of the variance in *s*_composite. It is therefore conceivable that the proxies for skills may in part capture an ideological effect of higher education. For example, much of the economic theory taught to university students during 1990s emphasized the efficiency of free markets over state intervention.[30]

To some extent we have already controlled for this possibility by including variables for people's assessment of their own level of information, as well as their support for parties on the left-right scale. If the highly educated consider themselves better informed about the costs of generous social spending, this is likely to show up in the variable measuring information. Likewise, those ideologically committed to a small welfare state are presumably more likely to support right parties. The fact that a large effect of skills persists after we control for these variables suggests that our conception of skills as assets is correct.

But there may still be unmeasured aspects of formal education that somehow confound the effects of our skill variable. One way to address this issue is simply to include general education as a separate variable. That way the effect of *s*_composite will only pick up the effects of specific skills. In this setup we would expect formal education to have the opposite effect than the specific skills variable, and the separating out of general skills will necessarily weaken the effect of the

[30] In terms of the formal model this can be captured by different assessments of the distortionary effects of taxation.

original variable if general education is indeed a measure of general skills. However, we can be certain that whatever effect remains of *s*, it cannot be attributed to general education.

The first column of Table 6 shows the results of re-estimating model (1) in Table 3, using formal education as a separate independent variable. Formal education has a strong *negative* effect on support for social protection. This is consistent with the skill asset argument. But more importantly, the parameter on the specific skills variable remains positive and statistically significant. Not surprisingly, the effect of *s* falls from 0.14 to 0.09, but this is still a very considerable impact. Even if we were to completely discount the effect of general education as a measure of general skills, therefore, the results lend unambiguous support to our argument.

[Table 6 about here]

Yet, we think it would be a mistake to treat general education as a proxy for unmeasured ideological effects, and we can support this claim with results for the "postmaterialist" spending index explained above. Surely, if highly educated individuals believe in the efficiency of free markets and the waste of government spending, they should also oppose public spending on the environment, culture, and the arts.[31] But the exact opposite is true as shown in column (2) of Table 6. People with high general education are much *more* likely to support government spending on these areas than others. Conversely, if we use our composite measure of specific skills (column 3), the effect of skills is reversed: specific skill workers want less "postmaterialist" spending even though they support more social spending. Evidently people prefer government spending in areas that are particularly conducive to their personal welfare. General skills workers demand little social protection but are enthusiastic consumers of a clean environment and state-subsidized culture. Specific skill workers are deeply concerned with social protection but not

[31] "It is true that "the environment" may be conceived as a collective good improving overall welfare (it is a little harder to argue this with respect to subsidies to the fine arts), but by the same token social protection may be conceived as welfare-improving insurance. The point is not that the highly educated are more informed about what is "good" and "bad" spending, we already control for this, but that they may have internalized a general aversion to government spending through their educational experience."

enthusiastic about state subsidization of environmental causes and the arts. There is no blanket support for, or opposition to, government spending among any particular group of workers.

Macro-level implications

There are two major macro-level implications of our argument. First, if skill profiles vary systematically across advanced economies, and if these differences are reflected in the political demands of the median voter, it suggests a new explanation of the welfare state driven politically by national patterns of skill specificity. Second, we show how such a theory can explain the observed inverse correlation between welfare spending and earnings inequality, an inequality which is difficult to reconcile with the rather clear statistical evidence that (a) increased inequality reduces the income of the median voter and (b) a fall in income increases the political demand for welfare state spending. We discuss each in turn.

A skill specificity-driven theory of the welfare state

A simple implication of our argument is that if skill profiles vary systematically across countries, political demand for social protection should also vary systematically across countries. Assuming that these demands find expression through the democratic political process -- as assumed in the median voter model, for example -- our micro-level argument can potentially account for a substantial portion of the macro-level variance in social protection across countries. In this section we explore this hypothesis, although we do not pretend to offer a conclusive test.

Our 11 countries divide rather cleanly into five with highly institutionalized vocational training systems, and six with poor vocational training systems (but good general education). In the first category of countries -- France, Italy, Germany, Norway, and Sweden -- a large percentage of an age cohort go through a longer vocational training (typically 3-5 years), and such training is geared towards developing specific skills: either for a particular firm, a particular industry, or some combination of the two. The share of an age cohort that goes through a

vocational training in these countries varies between 29 and 36 percent.[32]

Among the general skill countries -- Australia, Britain, Canada, Ireland, New Zealand, and the US -- only Britain and Ireland have anything that comes close to an institutionalized vocational training system. In both cases about 9 percent of an age cohort goes through vocational training. However it is widely recognized that genuine vocational training in Britain has been on the decline since the early 1970s (Wood 1997), and the so-called General National Vocational Qualifications system, introduced with the intention of providing higher-level vocational qualifications, is now seen as more akin to an intermediate general degree than to a vocational qualification. For the other general skills countries, initial vocational training tends to be weak and relatively short (2 years and less), and even counting short-term post-secondary degrees (such as the American junior college system) only between three and four percent of an age cohort acquire a vocational degree.

In order to examine whether these differences in skill systems are linked to differences in social protection, Figure 3 compares the two groups of countries on four indicators of social protection. The first is a composite measure of unemployment protection, the second is the OECD's measure of employment protection, the third is a measure of the degree of coordination in wage-setting (a key component of the wage protection system), while the fourth is government social transfers as a proportion of GDP. The last is closest the conception of social protection used in the model and in the analysis of public opinion. The numbers next to the country labels are the percentages of an age cohort receiving some kind of secondary or post-secondary vocational training.

What stands out from this comparison is that all the specific skills countries have notably higher levels of protection than any of the general skills countries. Using the mean of the four indicators as a proxy for the overall level of protection, a simple dummy variable for skill system explains 86 percent of the variance in protection. If we use the shares of an age cohort in some kind of vocational training as the independent variable, the explained variance is 84 percent. Since we know from the micro-level analysis that specific skill workers demand more social protection

[32] Based on UNESCO (1999) figures for the number of students in secondary vocational training or in short-term tertiary programs leading to vocationally oriented degrees (ISCED5).

than general skills workers, this precisely what we would expect to find.

[Figure 3 about here]

This conclusion is supported by data from a larger sample of 19 OECD countries.[33] The data are limited to the period 1980-1995 for which we have comparable figures for the share of an age cohort going through vocational training. This share serves as our macro-level proxy for the importance of specific skills in the labor force (the average *s*). The effect of relative income, the other variable in our theoretical model, is captured by a widely used OECD measure of (pre-tax) earnings inequality (d1/d9 ratios). If we control for per capita income (or mean income), an increase in inequality implies a reduction in the income of the median voter, assuming a usual right-skewed distribution of income. The effect of wage inequality therefore captures the importance of cross-national differences in the relative income of the median voter (*y*). Exactly as in the micro-level analysis the parameter for *s* is expected to be positive and the parameter for *y* to be negative.

In addition to the measures of relative income and skills, we control for the effects of variables that are widely hypothesized to affect social spending. Measures of the partisan composition of government and unionization rates are designed to capture the argument that a strong labor movement and left party control of government lead to more social spending (see, for example, Korpi 1983; 1989; and Huber and Stephens 2001). Trade openness, measured in terms of trade as a percentage of GDP, is designed to capture the argument that exposure to international competition causes labor market uncertainty and demand for compensating social spending (see Cameron 1978, Katzenstein 1985, and Rodrik 1998). Finally, GDP per capita tests "Wagner's Law," which says that demand for social services is income elastic.

Table 7 shows the results of a multiple regression using transfers as a share of GDP as the

[33] The countries are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, the Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, UK, and the US. We are missing data on some variables for Greece and the New Zealand, and have excluded them to facilitate the presentation.

dependent variable and inequality, vocational training, and the various controls as independent variables. Because the data are limited in time (a problem that is exacerbated by many missing observations on the inequality measure), and because spending is relatively stable and highly auto-correlated, there is little cross-time variation to be captured (almost 80 percent of the variance is cross-national). The data also turns out to be non-stationary (ñ>.99). The evidence presented here is therefore strictly cross-sectional.[34]

[Table 7 about here]

With this qualification, the findings for our theoretical variables are clearly supportive of our argument. Inequality has a small positive effect on spending, while vocational training has a large positive effect. Indeed, training is by far the most important variable in terms of both substantive impact and statistical significance. Thus, a one standard deviation increase in vocational training is associated with a similar increase in government transfers, and the vocational training variable accounts for between 35 (if entered as the last variable) and 65 percent (if entered as the first variable) of the total variance in spending. None of the other variables comes close in explanatory power, and none are in fact statistically significant.

A skill specificity explanation of the inverse correlation between welfare spending and inequality.

The finding of a positive, albeit small, effect of earnings inequality on spending highlights the empirical puzzle of a negative cross-national association between earnings inequality and social spending. In our data the bivariate correlation between the two variables is -0.42. How is it possible that egalitarian countries spend more on social protection when lower relative income is estimated to have a positive effect on support for social spending as well as actual spending? The answer is simple, and illuminating for the central role of skills: vocational training has a strong negative effect on earnings inequality *in addition to* its positive effect on spending. This effect is

[34] The problem of non-stationarity persists with the inclusion of country dummies (ñ>.99).

shown in the second column of Table 7. Specifically, a one standard deviation increase in the vocational training variable is associated with a .69 standard deviations reduction in inequality.

The simplest explanation for this effect is that a well developed vocational training system compresses the distribution of skills. The reason is that institutionalized vocational training systems offer opportunities to acquire valuable skills for young people who are academically disinclined, whereas in general skills systems these individuals tend to end up as low-paid unskilled workers. In so far as companies adjust their product market strategies to take advantage of the relative abundance of particular types of skills, early school leavers in specific skills systems will therefore tend to earn better wages than early school leavers in general skills systems.[35]

It is the *combination* of a negative effect of vocational training on inequality and its positive effect on redistributive spending that creates the negative bivariate cross-national association between inequality and spending. As illustrated in Figure 4, the negative bivariate relationship between inequality and spending is reversed once we control for training. It should be noted that although there is a negative *indirect* effect of vocational training on spending (through reduced inequality), this effect is easily outweighted by the positive direct effect. Taking into account both effects, the impact of a one standard deviation increase in training is to raise spending by .74 standard deviations.

[Figure 4 about here]

While the findings in this section cannot be considered definitive, the close correspondence between the micro and macro evidence strongly suggests that cross-national differences in skill systems are a source of significant cross-national variance in social spending. The structure of skills also affect the structure of earnings. Because expansive vocational training systems offer

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[35] The linkage between skills and wage compression is in some measure mediated by the collective wage bargaining system. Coordinated wage bargaining offer better wage protection as noted above, and this may encourage more young people to invest in specific skills. Conversely, once workers have invested in specific skills, they are more likely to join unions and support coordinated bargaining. For the dampening effect of centralized bargaining on wage inequality, see Rueda and Pontusson (2000) and Wallerstein (1999).

opportunities to acquire valuable skills for those who are academically disinclined, such systems are associated with a more even distribution of skills and income. Combining the two effects explains the puzzle of the cross-national association between equality and spending.

Conclusion

It is a well-known fact that a substantial portion of both national and personal income can be attributed to human capital, broadly conceived. It is therefore not surprising that the asset-specificity of this capital matters a great deal for the amount of social insurance demanded by individual workers. Like physical capital, human capital can be more or less mobile, and workers who have made heavy investments in asset-specific skills stand a greater risk of losing a substantial portion of their income than workers who have invested in portable skills. For this reason, specific skill workers have a greater incentive to support policies and institutions that protect their jobs and income.

Because social protection tends to benefit low-income people more than high-income people, position in the income distribution also divides public opinion. However, at any given level of income, workers with specific skills are more inclined to support high levels of protection than those with general skills. This helps us understand why countries with a highly developed vocational training systems, with a focus on cultivating specific skills, exhibit higher support for social protection *as well as* a more egalitarian distribution of income. Because developed vocational systems have a dampening effect on income inequality, while simultaneously creating high demand for social protection, equality and redistribution tend to go hand in hand. In the Meltzer-Richard model this is not possible because the pressure for redistribution is always greatest in countries with the most skewed income distribution.

If our argument is correct, then there is no necessary tension between redistributive social protection and competitiveness as argued by Alesina and Rodrik (1994). Because social protection serves as an insurance against loss of specific skill investments, social protection may well be a requisite for such investments in the first place. Firms that depend on a workforce with extensive specific skills to compete effectively in their chosen product markets can therefore benefit from high levels of social protection. Conversely, firms that depend primarily on workers

with general skills would be harmed by the costs of social protection without benefitting from the supply-side effects of higher protection. Whether social protection undermines competitiveness therefore depends entirely on the position of countries in the international division of labor.

Finally, our model points to an important source of cross-time variance in support for social protection: unanticipated shocks to the occupational structure. If workers have sunk investments in skills that are not fully transferable, an increase in the risk of having to move across a skill boundary in the economy raises the level of demand for social insurance. This helps us understand why the dramatic decline of industrial employment in many countries over the past three decades is a very good predictor of welfare state expansion (Iversen and Cusack 2000). More generally, changes over time in the exposure to risks, changes in the training system, and technologically induced changes in the international division of labor all affect the political demand for social protection. Modeling these dynamics, and testing them empirically, are important tasks for future research.

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Appendix A:
Mathematical proofs.

***Derivation of results (12) and (13) in section (iii)**:*

The choice of the optimal *R* requires that:

$$V_R \geq 0 \Leftrightarrow \; b \cdot u'(\bar{g}) \cdot 2g / w = g \cdot u'(R)$$

Totally differentiating both sides we get:

$$\frac{dR}{dg} = \frac{\frac{2b}{w} \cdot [\bar{g} \cdot u''(\bar{g}) + u'(\bar{g})]}{b \cdot \left(\frac{2g}{w}\right)^2 \cdot u''(\bar{g}) + g \cdot u''(R)}$$

Since the denominator is negative,

$$\frac{dR}{dg} > 0 \;\; iff \;\; [\bar{g}u''(\bar{g}) + u'(\bar{g})] < 0$$

which implies

$$RRA(\bar{g}) \equiv -\frac{\bar{g}u''(\bar{g})}{u'(\bar{g})} > 1$$

where *RRA(x)* is the Arrow-Pratt definition of Relative Risk Aversion defined at *c=x*. The inequality conditions specified in (12) and (13) follow directly.

Proof for Result I in section (iv):

(i) Note first that *t*=1 maximizes *t*/(1+*t*) when $0 \leq t \leq 1$. Also, if *t*=1, *R*=*w*/2.

(ii) From (6) the necessary condition for optimal *R* is

$$(1A) \qquad a \cdot u'(\bar{sg}) \cdot (1 - \frac{2sg}{w}) + b \cdot u'(\bar{g}) \cdot (1 - \frac{2g}{w}) + g \cdot u'(R) \geq 0$$

If $R = w/2$, $\bar{sg} = \bar{g} = R$; hence the maximum combination of *sg* and *g* at which *R*=*w*/2, assuming it exists, requires that this condition holds with equality and that $u'(\bar{sg}) = u'(\bar{g}) = u'(R)$. These conditions imply directly that $a \cdot sg + b \cdot g = (a + b + g) \cdot \frac{w}{2}$ $= y = \frac{w}{2}$.

Proof for Results II and III of model (iv):

The necessary condition for optimal choice of R is $V_R(R, s, g) = 0$. This is given by 1A above.

Totally differentiating V_R gives:

$$
\begin{aligned}
&a \cdot \left[u''(\bar{s}\bar{g}) \cdot \left(\frac{2sg}{w} - 1 \right) \cdot \left(1 - \frac{2R}{w} \right) \cdot g + u'(\bar{s}\bar{g}) \cdot g \cdot \frac{2}{w} \right] \cdot ds \\
&+ a \cdot \left[u''(\bar{s}\bar{g}) \cdot \left(\frac{2sg}{w} - 1 \right) \cdot \left(1 - \frac{2R}{w} \right) \cdot s + u'(\bar{s}\bar{g}) \cdot s \cdot \frac{2}{w} \right] \cdot dg \\
&+ b \cdot \left[u''(\bar{g}) \cdot \left(\frac{2g}{w} - 1 \right) \cdot \left(1 - \frac{2R}{w} \right) + u'(\bar{g}) \cdot \frac{2}{w} \right] \cdot dg \\
&= \left\{ a \cdot u''(\bar{s}\bar{g}) \cdot \left(\frac{2sg}{w} - 1 \right)^2 + b \cdot u''(\bar{g}) \cdot \left(\frac{2g}{w} - 1 \right)^2 + g \cdot u''(R) \right\} \cdot dR
\end{aligned}
\tag{2A}
$$

Note: (i) The term in curly brackets on the RHS, which we will call B, is negative. (ii) We can write $\left(\bar{s}\bar{g} - w/2\right) \cdot (1 - 2R/w) = \bar{s}\bar{g} - \bar{w}/2$. And (iii):

$$
\begin{aligned}
&\left[u''(\bar{s}\bar{g}) \cdot \left(\bar{s}\bar{g} - \bar{w}/2 \right) + u'(\bar{s}\bar{g}) \right] \\
&= u'(\bar{s}\bar{g}) \cdot \left[1 - RRA \cdot \frac{\bar{s}\bar{g} - \bar{w}/2}{\bar{s}\bar{g}} \right] \equiv u'(\bar{s}\bar{g}) \cdot L(\bar{s}\bar{g})
\end{aligned}
\tag{3A}
$$

So (2A) can be written:

$$
\begin{aligned}
&u'(\bar{s}\bar{g}) \cdot L(\bar{s}\bar{g}) \cdot a \cdot g \cdot ds + u'(\bar{s}\bar{g}) \cdot L(\bar{s}\bar{g}) \cdot a \cdot s \cdot dg \\
&+ u'(\bar{g}) \cdot L(\bar{g}) \cdot b\, dg = (w/2) \cdot B \cdot dR
\end{aligned}
\tag{4A}
$$

Since $dy = a \cdot g \cdot ds + a \cdot s \cdot dg + b \cdot dg$, we can further rewrite (4A) as

$$
\begin{aligned}
dR = &\frac{2a \cdot b \cdot g}{wB} \cdot \left[\frac{u'(\bar{s}\bar{g}) \cdot L(\bar{s}\bar{g}) - u'(\bar{g}) \cdot L(\bar{g})}{a\,s + b} \right] \cdot ds \\
&+ \frac{2}{wB} \cdot \left[\frac{u'(\bar{s}\bar{g}) \cdot L(\bar{s}\bar{g}) \cdot a\,s + u'(\bar{g}) \cdot L(\bar{g}) \cdot b}{a\,s + b} \right] \cdot dy
\end{aligned}
\tag{5A}
$$

To prove **Results III** and **IV**, note that in terms of (5A) $\partial R / \partial y = dR / dy$ and $\partial R / \partial s = dR / ds$. We show first that $L(\bar{s}\bar{g}) < L(\bar{g})$. From the definition in (3A), this follows if $s > 1$ – as is the case apart from purely general skills – and if RRA > 0. **Result III** is that

$\text{sgn}\,\partial R/\partial y < 0$ *if* $RRA < \bar{s}\bar{g}/(\bar{s}\bar{g} - \bar{w}/2)$. Since $B < 0$, $L(\bar{s}\bar{g}) < L(\bar{g})$ and $u'(x) > 0$, this follows from (5A) if $L(\bar{s}\bar{g}) > 0$. This requires that $RRA < \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2}$. This is a sufficient condition: a necessary and sufficient condition is that the numerator of the second term in square brackets on the RHS of (5A) is positive.

Result IV is that $\text{sgn}\,\partial R/\partial s > 0$. Since $B < 0$, this requires that the numerator in the first square bracket on the RHS of (5A) is negative. Since $u'(\bar{s}\bar{g}) < u'(\bar{g})$ from diminishing marginal utility, a sufficient condition is that $L(\bar{s}\bar{g}) < L(\bar{g})$, which is true so long as RRA > 0 and s > 1. So **Result IV** follows from the existence of risk aversion and specific skills.

Appendix B:
Deriving the estimating equation

In this appendix we show that the estimating equation that we use:

(6A) $$R = k + a \cdot y + b \cdot s$$

is equal to

(7A) $$R = k + \frac{\partial R}{\partial y} \cdot y + \frac{\partial R}{\partial s} \cdot s$$

where (7A) is a first-order Taylor expansion of $V_R(R, s, g) = 0$ and $y = a \cdot s \cdot g + b \cdot g$ evaluated around $(R, s, g) = (\tilde{R}, \tilde{s}, \tilde{g}) \equiv \tilde{x}$

Proof: The first-order Taylor expansion of V_R is given by:

(8A) $$R = K + \frac{V_{R,s}}{V_{R,R}} s + \frac{V_{R,g}}{V_{R,R}} g$$

In terms of (5A):

(9A.1) $$\frac{V_{R,s}(\tilde{x})}{V_{R,R}(\tilde{x})} = \frac{u'(\bar{s}\bar{g}).L(\bar{s}\bar{g}).a.\tilde{g}}{(w/2).B}$$

and

(9A.2) $$\frac{V_{R,g}(\tilde{x})}{V_{R,R}(\tilde{x})} = \frac{u'(\bar{s}\bar{g}).L(\bar{s}\bar{g}).a.\tilde{s} + u'(\bar{g}).L(\bar{g}).b}{(w/2).B}$$

The first-order Taylor expansion of y is:

(10A) $$y = k(\tilde{x}) + \left[a.\tilde{s} + b\right].g + \left[a.\tilde{g}\right].s$$

Rewrite (10A): $g = \frac{y - k(\tilde{x}) - a\tilde{g}}{a\tilde{s} + b}$ and substitute into (8A), using (9A.1) and (9A.2). This yields (7A).

Appendix C:
Statistical appendix

A problem arises in our use of *s_3* and *s_4* as explanatory variables. (Since it is the same in both cases we will simply refer to *s*.) Because the question used as the basis for *s* was asked only in the 1997 survey, whereas all the questions about spending were asked only in the 1996 survey, it was necessary to "translate" the 1997 information on *s* so it could be used in the 1996 survey. For this purpose we calculated averages for *s* at the 3-digit ISCO-88 level in the 1997, and then assigned these values to individuals in the 1996 survey based on their 3-digit ISCO classification in that survey.

We show here that the estimated coefficient of *b* has an approximate small sample bias which biases down the estimated coefficient towards zero if $b > 0$, and biases it upwards towards zero if $b < 0$.

The structural model is

$$\text{(C.1)}\quad R_{i,j}^{96} = k + a.y_{i,j}^{96} + b.s_{i,j}^{96} + e_{i,j}^{96}$$

where each observation is drawn from the 1996 survey and where *i* indexes the individual and *j* the ISCO 3-digit level. We do not have data on $s^{96}_{i,j}$. Assume $s^{96}_{i,j}$ is generated by the process

$$\text{(C.2)}\quad s_{i,j}^{96} = s_j + h_{i,j}^{96}$$

where s_j is exogenous. s_j itself is unobservable but we have data from the 1997 survey generated by the same process:

$$\text{(C.3)}\quad \begin{aligned} & s_{i,j}^{97} = s_j + h_{i,j}^{97} \\ & Eh_{i,j}^{9x} = 0 \quad \forall i, j, x \\ & Eh_{i,j}^{96}.h_{r,j}^{97} = 0 = Ehe \quad \forall i, r, j \\ & Eh^2 = s_h^2;\ Ee^2 = s_e^2 \end{aligned}$$

We now run the regression

$$\text{(C.4)}\quad \begin{aligned} & R_{i,j}^{96} = k + ay_{i,j}^{96} + b\bar{s}_j + e_{i,j}^{96} + u_{i,j}^{96} \\ & \text{where}\quad \bar{s}_j \equiv \frac{\sum s_{i,j}^{97}}{N_j} \\ & \text{and}\quad u_{i,j}^{96} \equiv b\left[s_{i,j}^{96} - \bar{s}_j\right] \end{aligned}$$

where N_j is the number of individuals in ISCO category j in the 97 survey. From (C.2, 3)

$$(C.5) \qquad u_{i,j}^{96} = h_{i,j}^{96} - \frac{\sum h_{i,j}^{97}}{N_j}$$

We can simplify the exposition considerably by assuming that there is no correlation between $\bar{s}$ and y. This implies:

$$E\hat{a} = a$$

$$(C.6) \qquad E\hat{b} = b\left(1 - E\frac{\sum_j \bar{s}_j.\bar{h}_j}{\sum_j \bar{s}_j^2}\right)$$

Making the appropriate probability limit assumptions it is not difficult to show that $\hat{b}$ is a consistent estimator of b. We can get a better insight from the approximation

$$(C.7) \qquad E\frac{\sum \bar{s}\bar{h}}{\sum \bar{s}^2} \approx \frac{E\sum \bar{s}\bar{h}}{E\sum \bar{s}^2}$$

(which holds so long as the sample is not very small). Let there be J ISCO categories and assume for convenience that $N_j = N \ \forall j$. Then we have:

$$(C.8) \qquad E\hat{b} = b\left(1 - \frac{s_h^2}{s_h^2 + N.\frac{\sum s_j^2}{J}}\right)$$

Since J is constant, (C.8) tells us first that as N increases the approximate bias goes to zero. Second and more important, it implies that for a small sample

$$(C.9) \qquad \begin{aligned} &b < E\hat{b} < 0 \quad if \quad b < 0; \\ &b > E\hat{b} > 0 \quad if \quad b > 0. \end{aligned}$$

Finally, it implies that

$$(C.10) \qquad If \quad b = 0 \Rightarrow E\hat{b} = 0$$

and the standard significance tests hold.

Tables and Figures

Table 1. Principal factor analysis of six spending variables (*N*=13,557)

	Factor (varimax rotation)	
Policy area:	Social spending	Postmaterialist spending
Pensions	**0.64**	0.07
Health care	**0.54**	0.14
Unemployment	**0.52**	0.23
Subsidies	**0.46**	0.04
Culture and the arts	0.09	**0.48**
Environment	0.15	**0.48**
Eigenvalues	1.4	0.4

Table 2. Summary of independent skill variables.

Variable name:	Definition:	Inter-correlations				Comment:
		*s*_1	*s*_2	*s*_3	*s*_4	
*s*_1	(Share of ISCO-88 level 4 groups)/ (share of labor force) divided by ISCO level of skills[1)]	1				
*s*_2	(Share of ISCO-88 level 4 groups)/ (share of labor force) divided by level of general education[1)]	.82	1			
*s*_3	Response to question about difficulty of finding an acceptable job.	.16 (.38)	.15 (.53)	1		Not clear whether this is a measure of absolute or relative skills
*s*_4	*s*_3 divided by level of general education	.33 (.66)	.55 (.59)	.62	1	Assumes that *s*_3 measures absolute skills (though *s*_4 will always be a relative measure)

[1)] Shares are calculated at both the first and second ISCO-88 level and then averaged.

Table 3. Support for social spending among the publics of 10 OECD countries, 1996 (t-scores in parentheses).

	Dependent variable: Support for social spending[1)]					
	(1)	(2)[2)]	(3)[2)]	(4)	(5)	(6)[3)]
Income	-0.0023**	-0.0031**	-0.0026**	-0.0025**	-0.0026**	-0.0026**
	(-17.33)	(-18.84)	(-15.99)	(-19.20)	(-19.54)	(-16.00)
*s*_composite	0.146**					0.134**
	(19.72)					(18.53)
*s*_1		0.095**				
		(15.39)				
*s*_2			0.097**			
			(16.38)			
*s*_3				0.072**		
				(8.71)		
*s*_4					0.132**	
					(16.43)	
Age	0.002**					0.002**
	(5.21)					(4.09)
Gender (female)	0.14**					0.14**
	(12.02)					(10.93)
Union membership	-					0.12**
						(8.14)
Part-time empl.	-0.02					-0.02
	(-1.05)					(-1.07)
Unemployed	0.19**					0.21**
	(7.15)					(7.56)
Non-employed	-0.05**					-0.03
	(-3.38)					(-1.66)
Self-employed	-0.15**					-0.12**
	(-8.11)					(-6.76)
Informed	0.03**					0.03**
	(5.03)					(4.87)
L-R party support	-0.03**					-0.03**
	(-11.12)					(-8.93)
Adj. R-squared	0.21	0.20	0.20	0.18	0.20	0.22
N	14,101	10,956	10,956	14,101	14,101	11,950

Key: * significant at the .05 level; ** significant at the .01 level.

Notes: [1)] All regressions included a full set of country dummies (not shown); [2)] excludes Australia, Ireland, and Italy for which data are not available; [3)] excludes Australia for which union membership data are not available.

Table 4. Estimates of the magnitude of the effects of independent variables.

	Proportion of explained variance[1)]		Impact of a one std. change[4)]	
	Lower bound[2)]	Upper bound[3)]	95% confidence interval	
Income	16	53	-0.14	-0.12
*s*_composite	19	37	0.13	0.14
Age	1	1	0.02	0.03
Gender (female)	6	17	0.06	0.07
Union membership	1	4	0.05	0.06
Part-time employment	0	0	-0.01	-0.00
Unemployed	3	8	0.04	0.05
Non-employed	0	8	-0.02	-0.01
Self-employed	3	8	-0.05	-0.04
Informed	2	8	0.02	0.04
L-R party support	4	5	-0.06	-0.05
Income *and s*_comp	48	73	0.25	0.28
All controls combined	27	52	0.24	0.32

Notes: [1)] Increase in explained variance by each variable as proportion of the total explained variance of all variables. [2)] Increase in explained variance (compared to model with only country dummies) when each variable is included as the *first* variable. [3)] Increase in explained variance when a variable is included as the *last* variable. [4)] The change in support for social spending as a result of a one standard deviation increase in each of the independent variables. The last two rows assume change in the independent variables that raise support for spending (and take into account that some combinations of the employment variables are impossible).

Table 5. Income, skills, and support for social spending in 11 OECD countries (t-scores in parentheses).[1)]

	Income[2)]	***s*_comp**	*s*_1	*s*_2	*s*_3	*s*_4	*N*
Australia	-0.0020**	**0.103****	0.076**	0.078**	n.a[3)]	n.a	2151
	(-6.68)	**(5.92)**	(4.68)	(4.36)			
Britain	-0.0020**	**0.131****	0.066**	0.079**	0.086**	0.133**	989
	(-4.57)	**(5.10)**	(3.73)	(4.39)	(3.27)	(4.66)	
Canada	-0.0039**	**0.115****	0.067*	0.091**	0.064**	0.131**	1182
	(-6.85)	**(4.17)**	(2.42)	(3.11)	(2.60)	(4.51)	
France	-0.0037**	**0.123****	0.098**	0.085**	0.067**	0.079**	1312
	(-8.53)	**(5.45)**	(4.20)	(4.40)	(2.63)	(4.77)	
Germany	-0.0022**	**0.141****	0.106**	0.092**	0.076**	0.143**	2361
	(-5.04)	**(8.24)**	(6.85)	(6.28)	(3.26)	(7.67)	
Ireland	-0.0020**	**0.075****	0.060**	0.080**	n.a	n.a	994
	(-3.85)	**(4.10)**	(3.14)	(4.51)			
Italy	-0.0014	**0.065****	0.065*	0.059**	n.a	n.a	983
	(-1.32)	**(2.81)**	(2.57)	(2.75)			
Norway	-0.0016**	**0.147****	0.089**	0.111**	0.049**	0.166**	1344
	(-4.23)	**(7.48)**	(5.97)	(6.40)	(2.70)	(7.45)	
New Zealand	-0.0034**	**0.107****	0.086**	0.064**	0.090**	0.085**	1198
	(-6.93)	**(5.12)**	(4.57)	(2.82)	(2.77)	(3.74)	
Sweden	-0.0042**	**0.160****	0.083**	0.093**	0.088**	0.187**	1238
	(-5.89)	**(7.51)**	(4.51)	(5.16)	(4.08)	(8.09)	
United States	-0.0016**	**0.170****	0.123**	0.130**	0.031	0.198**	1332
	(-2.87)	**(6.19)**	(6.49)	(5.56)	(1.23)	(5.03)	

Key: * significant at the .05 level; ** significant at the .01 level.

Notes: [1)]All regressions included the same set of controls as in Table 3; [2)] The effect of income is only shown for *s*_comp, but varies little across the four measures of *s*; [3)] n.a.: data not available to estimate this parameter.

Table 6. Formal education and support for two types of spending in 10 OECD countries, 1996 (t-scores in parentheses).

	Support for social spending	Support for postmaterialist spending	
Income	-0.0019**	-0.0003	0.0003
	(-14.09)	(-1.88)	(2.07)
*s*_comp	0.088**	-	-0.067**
	(10.25)		(-7.35)
Formal education	-0.070**	0.100**	-
	(-13.34)	(18.36)	
Age	0.00*	-0.004**	-0.006**
	(2.24)	(-9.80)	(-13.41)
Gender (female)	0.14**	0.07**	0.07**
	(11.53)	(5.03)	(4.75)
Part-time employment	-0.02	0.08**	0.08**
	(-0.92)	(3.53)	(3.80)
Unemployed	0.20**	0.07	0.07
	(7.39)	(1.99)	(2.05)
Non-employed	-0.05**	0.06**	0.07**
	(-2.93)	(3.19)	(3.81)
Self-employed	-0.16**	0.02	0.01
	(-8.41)	(0.83)	(0.45)
Informed	0.02**	-0.05**	-0.06**
	(3.79)	(-7.54)	(-9.17)
L-R party support	-0.03**	-0.05**	-0.04**
	(-11.05)	(-12.54)	(-12.25)
Adj. R-squared	0.22	0.09	0.07
N	14,101	14,101	14,101

Key: * significant at the .05 level; ** significant at the .01 level.

Note: Regressions included a full set of country dummies.

Table 7. The determinants of government transfers in 19 OECD countries, 1980-95 (t-scores in parentheses).

	Dependent variable:	
	Government transfers as percent of GDP[1)8)]	Earnings inequality[2)8)]
Constant	4.06	4.38
	(1.86)	(3.07)
Inequality[2)8)]	0.44	-
	(1.38)	
Vocational training[3)]	0.019**	-0.012**
	(3.97)	(-4.56)
Left government[4)]	0.10	0.03
	(1.08)	(0.43)
Unionization[5)]	0.001	-0.004*
	(0.59)	(-2.76)
Trade openness[6)]	0.06	0.04
	(1.53)	(1.19)
GDP per capita [7)8)]	-0.28	-0.34*
	(-1.26)	(-2.00)
Adj. R-squared	0.68	0.73
N	19	19

Key: * significant at the .05 level; ** significant at the .01 level.

Notes: [1)]All government payments to the civilian household sector (including social security transfers, government grants, public employee pensions, and transfers to non-profit institutions serving the household sector) as a percent of GDP. *Sources*: Cusack (1991) and OECD, *National Accounts* (various years). [2)] The earnings of worker in the top decile of the earnings distribution relative to a worker in the bottom decile of the earnings distribution. *Source*: OECD, *Electronic Data Base on Wage Dispersion* (undated). [3)] The share of an age cohort in either secondary or post-secondary (ISCED5) vocational training. *Source*: UNESCO (1999). [4)] This is an index of the partisan left-right "center of gravity" developed by Cusack (1997). The index varies from 0 (extreme right) to 4 (extreme left). [5)] Union density rates. *Sources*: Ebbinghaus and Visser (2000). [6)] Total exports and imports of goods and services as percentage of GDP. Source: OECD, *National Accounts* (various years). [7)] GDP per capita. *Source*: Summers and Heston (1999). [8)] Log-transformed to improve fit.

Figure 1: The three states of an individual in the labor market.

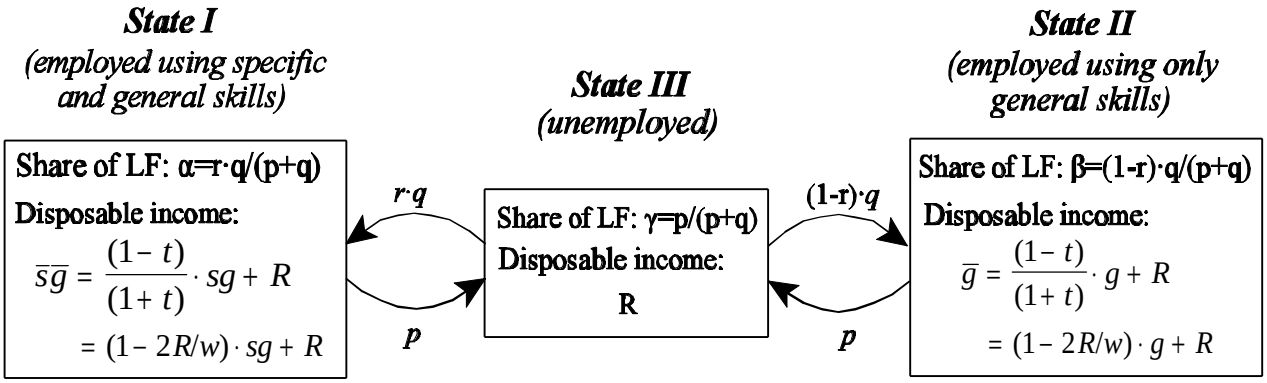
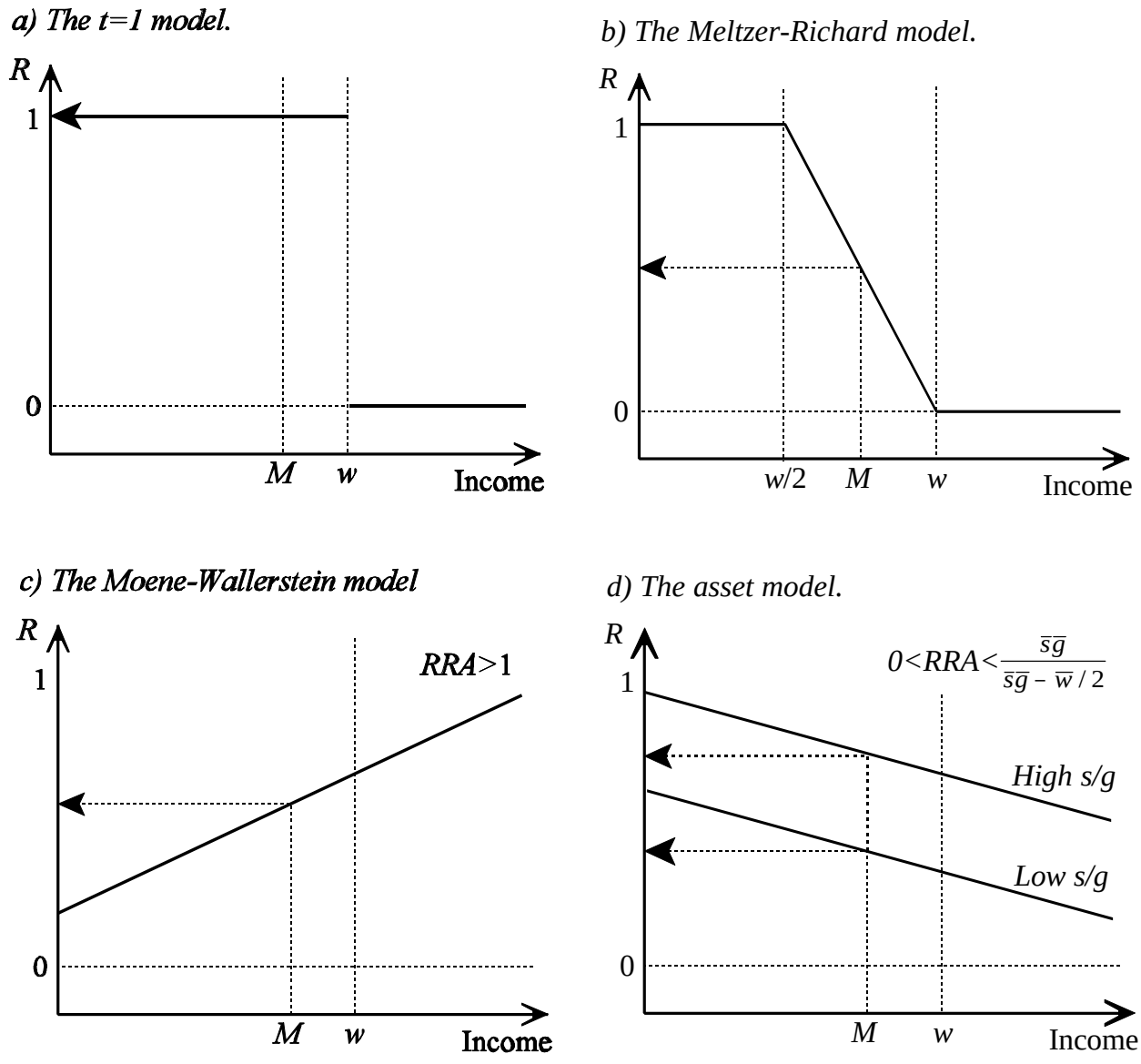
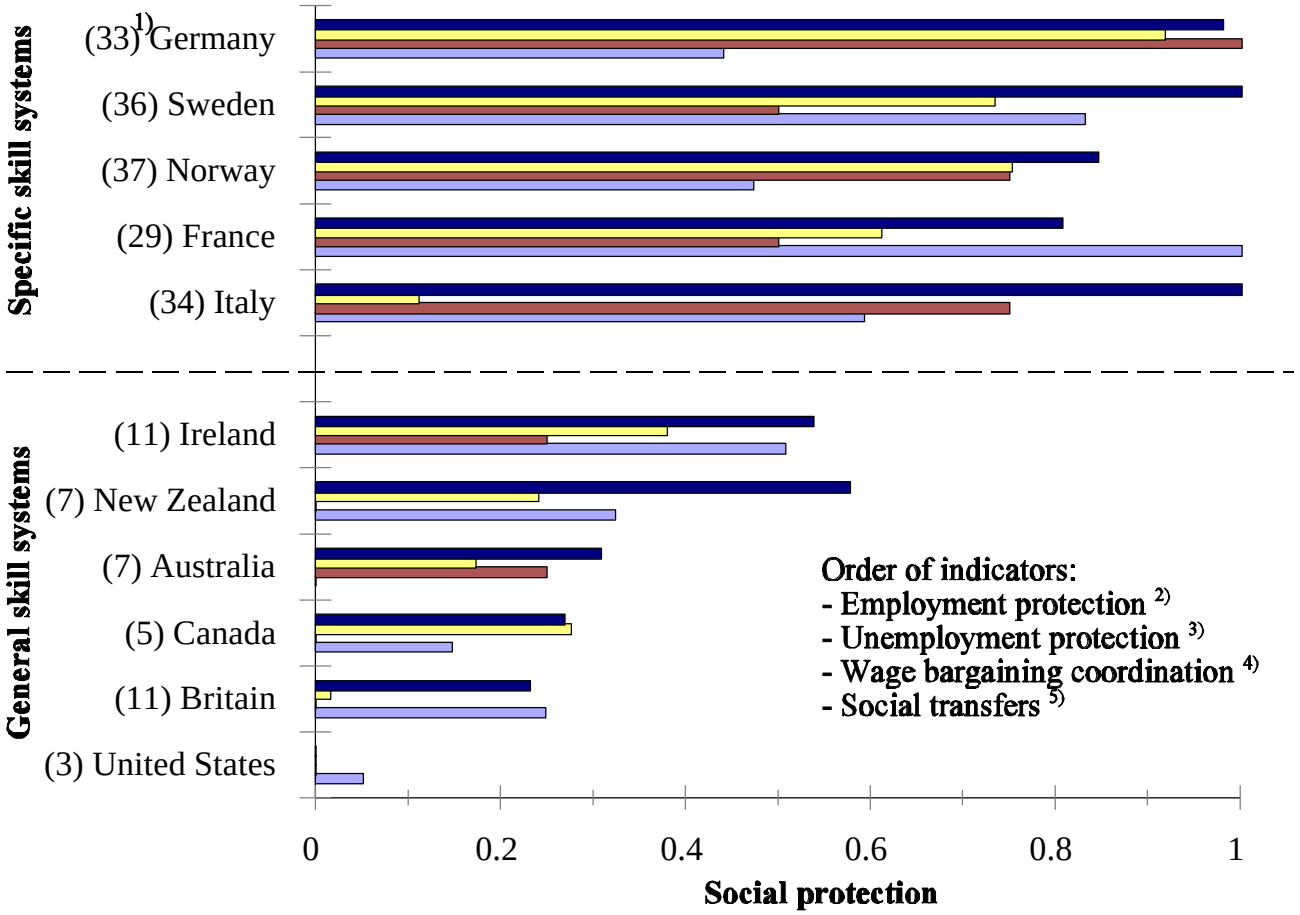


Figure 2. Four models of social policy preferences.



Note: Arrows indicate preferred level of social protection by the median voter.

Figure 3. Skill systems and social protection.



[1] The share of an age cohort in either secondary or post-secondary (ISCED5) vocational training. *Source*: UNESCO (1999).

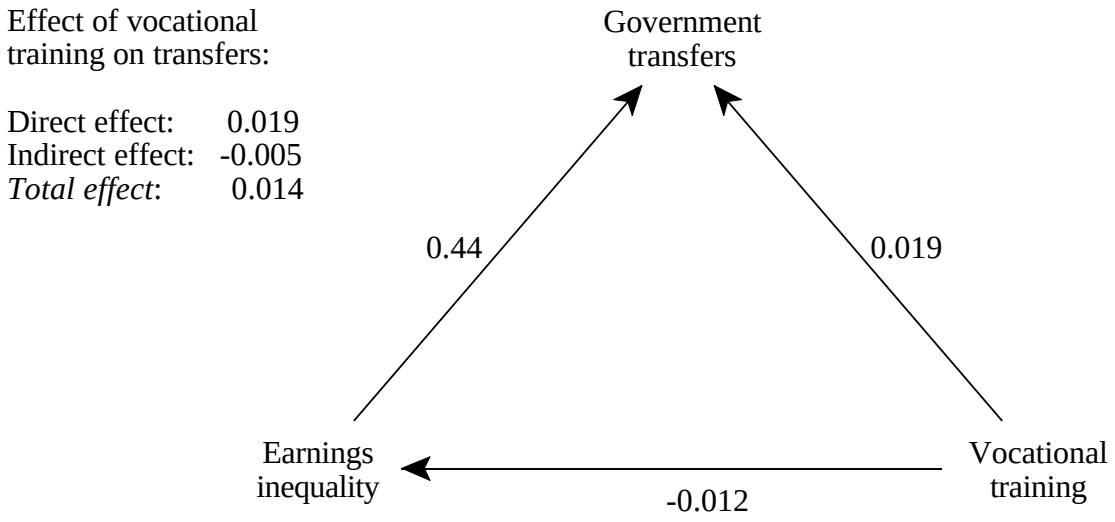
[2] OECD's index of employment protection based on the "restrictiveness" of individual hiring and firing rules, as well as collective dismissal rules. *Source*: *OECD Employment Outlook* (1999).

[3] Average of three indicators: a) net unemployment replacement rates for a 40 year old representative worker; b) The share of GDP paid in unemployment benefits as a percent of the share of the unemployed in the total population; and c) Index that measures the restrictiveness of the definition of a "suitable job" in the administration of benefits to the unemployed. *Sources*: See Estevez et al. 1999.

[4] Degree of coordination in wage-setting across firms (1994). *Source*: *OECD Employment Outlook* (1997, p. 71).

[5] All government payments to the civilian household sector, including social security transfers, government grants, public employee pensions, and transfers to non-profit institutions serving the household sector as a percent of GDP (1980-95). *Sources*: Cusack (1991) and OECD, *National Accounts, Part II: Detailed Tables* (Paris: OECD, various years).

Figure 4. The causal relationship between vocational training, earnings equality, and government transfers.



Note: The effects of vocational raining and of earnings inequality are the estimated parameters from the regression results in Table 7.